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On the Purchasing Power Parity for Latin-American Countries*

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Abstract

The purpose of this paper is to test the hypothesis of long-run purchasing power parity (PPP) for all Latin American countries. These countries share similar economic history and contagious effects from currency crises, which might lead to co-movements in their real exchange rates. New time series unit root tests found evidence of PPP for the vast majority of countries. In the panel data framework, tests for the null of unit root, null of stationarity, and unit root under multiple structural breaks indicated stationary real exchange rates. Thus, there is convincing evidence that PPP holds for Latin-American countries in the post-1980 period.

JEL Classification: C12; C32; E43; F31.

Key Words: Purchasing Power Parity; Panel Data; Unit Root Tests; Latin America.

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1 Introduction

Long-run Purchasing Power Parity (PPP) is a corner-stone of many theoretical models in international economics. One way of interpreting the PPP doctrine is that real exchange rates should be mean-reverting, meaning that in response to any shock the real exchange rate must eventually return to its PPP defined level. This is a useful interpretation because it is empirically testable by unit roots tests. Empirical studies along these lines, however, rarely rejected a unit root in real exchange rates when using traditional augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, widely recognized as suffering from low power and size distortions.

Recent developments in time series and panel data econometrics have provided better tests to look for evidence on the PPP hypothesis. New tests proposed by Elliott, Rothenberg, and Stock (1996) and Ng and Perron (2001) display considerable gains in power and size compared to the traditional ADF and PP tests. In the panel data framework, test by Nyblom and Harvey (2000) on the null of stationarity works as complement to panel tests based on the null of unit root, which frequently over-reject the null when a few individuals in the panel are stationary. Occurrence of structural breaks also affect power of unit root tests. So, a successful PPP testing strategy should apply tests that allow for shifts in real exchange rates.

While there is great amount of empirical work testing for the PPP hypothesis in developed countries, much less effort has been spent to test it in developing countries. Specifically, there is a lack of evidence for the Latin-American countries taken as

whole. These countries share important similarities in their economic history, which might lead to co-movements in their real exchange rates. In addition, as argued by Calvo et al (1993) and Calvo and Reinhardt (1996), currency crises that spread over the region showed contagious effects that have led to narrow time dispersion in structural breaks¹. In the post-Bretton Woods period, for instance, they have faced high inflation, low average of economic growth, and successive economic stabilization plans with frequent intervention in the exchange rate regimes. These common features require that a pooled real exchange rate should be considered when testing for the PPP in the region.

The goal of this paper is to test the hypothesis of long-run PPP for all Latin-American countries in the post-1980 period using both time series and panel data unit root tests. We apply new time series tests, with good size and power, and recent panel data unit root tests. In both cases, tests that allow for structural breaks are performed. The possibility of non-linearity in real exchange rates is also considered. Our major contribution is to show evidence of long-run PPP under both new time series unit root tests and panel data tests for pooled real exchange rates of Latin-American countries. Our panel data results are not sensible to the null hypothesis of a unit root or multiple structural breaks in real exchange rates.

Taylor and Taylor (2004) present an excellent survey of the related literature and conclude that there has been a general acceptance of the empirical evidence on

¹See Eichengreen et al (1993) for a discussion on how to test for contagious effects.

the long-run PPP. Though the puzzle still continues on the volatility of short-run exchange rates and the long-run effect of adjustment through the PPP. Favorable evidence of long-run PPP is provided, for instance, by Taylor (2002). An important contribution of Taylor's work is to construct real exchange rate data for over 100 years for 20 countries². Based mainly on the results of the DF-GLS test, due to Elliott, Rothenberg, and Stock (1996), Taylor concludes that PPP has held in the long-run over his secular sample.

Moving to panel data unit root tests, one increases power to reject the null. However, power is still an issue if the time period entails breaks in the series. Papell (2002) proposes a panel unit root test that allows for three breaks chosen endogenously in the changing growth model of Perron (1989). He applied the test to a panel of 21 industrialized countries from 1973 to 1996 to model structural breaks in the 80's, where a significant depreciation took place after a large appreciation of the dollar. He was successful in rejecting a unit root in panels up to 15 typical countries.

For developing countries, Alba and Park (2003) analyzed a sample of 65 countries during the current floating period, from 1976 to 1999. They partitioned the data in two 10-year periods and organized the data according to country characteristics. Yet, by applying traditional tests, they found only limited support for long-run PPP. This result, however, might be biased because traditional unit root tests are severely affected by structural changes, as the ones that struck the developing-country real

²Taylor's (2002) sample include 18 developed countries plus Argentina and Brazil in the period from 1892 to 1996.

exchange rates during the period.

The paper is organized as follows. Next section briefly discuss the PPP theory and testing approach. Section 3 presents the time series unit root tests. Section 4 displays the panel data unit root tests. The results are reported and analyzed in section 5. Finally, section 6 is dedicated to concluding remarks.

2 Theoretical Background

The absolute version of the PPP states that national price levels should be equal when converted to a common currency and it is usually expressed as:

$$P_t = \epsilon_t P_t^* \tag{1}$$

where P is the home-country price level, P^* is the foreign-country price level, and ϵ is the nominal exchange rate. Equation (1), however, does not find favorable empirical evidence. Common reasons used to justify failure of the absolute PPP include existence of transportation costs and commercial barriers, presence of non-tradable goods in the price indexes, and difference in preferences across countries.

Because of the strong restriction imposed by (1), according to which the real exchange rate is constant and equals to one, empirical evidence of the PPP has focused on a weaker version, which states that the (log) real exchange rate obtained from (1) is stationary. In this case, deviations from the PPP are temporary and mean

reverting. For a single-country, one can test this weak version by:

$$\log \left(\frac{\epsilon_t P_t^*}{P_t} \right) \equiv q_t = \alpha + \xi_t \quad (2)$$

where, if the PPP holds, q_t must be stationary. From (2), a natural way to test for PPP is through unit root tests.

In a cross-country environment, equation (2) can be rewritten as:

$$\log \left(\frac{\epsilon_{it} P_{it}^*}{P_{it}} \right) \equiv q_{it} = \alpha_i + \xi_{it} \quad (3)$$

where $i = 1, 2, \dots, N$ countries, $t = 1, 2, \dots, T$ time periods. The compound error term, ξ_{it} , is assumed to be *i.i.d.* across i and over t . The PPP can be tested in (3) by panel data unit root tests, which have better power than the time series ones.

3 Time Series Unit Root Tests

We start by testing for a unit root in real exchange rates using the familiar ADF test, due to Dickey and Fuller (1979, 1981) and Said and Dickey (1984), and the Z_α (or PP) test, due to Phillips (1987) and Phillips and Perron (1988). Critical values for the τ -distribution from a large set of simulations are given by Mackinnon (1991).

It is well known, however, that the previous tests display serious distortions in power and size. Improvements in the test procedure have been proposed by Perron

and Ng (1996), Elliott, Rothenberg and Stock (1996), and Ng and Perron (2001).

In the way of the new tests, Elliott, Rothenberg and Stock (1996) show that OLS detrending is inefficient when there is high persistency in the data and suggest to use GLS detrended data. Let \tilde{q}_t be the GLS detrended version of q_t . Then, $\tilde{q}_t = q_t - \hat{\alpha}' z_t$, where the GLS coefficient $\hat{\alpha}$ is obtained as follows. Let $q_t^d = q_t - \bar{\alpha} q_{t-1}$ for $t = 2, 3, \dots, T$ and $q_1^d = q_1$. Define z_t^d in the same way. Then, we obtain $\hat{\alpha}$ in an OLS regression of q_t^d on z_t^d . The value of $\bar{\alpha}$ is given by $\bar{\alpha} = 1 + \bar{c}/T$, where \bar{c} depends on the deterministic terms included in z_t . As stated by Elliott, Rothenberg and Stock (1996), one should set $\bar{c} = -7$ if $z_t = \{1\}$ and $\bar{c} = -13.5$ if $z_t = \{1, t\}$.

The ADF^{GLS} test is given by t-statistic on the null hypothesis that $\beta = 0$ at:

$$\Delta \tilde{q}_t = \beta \tilde{q}_{t-1} + \sum_{j=1}^k \gamma_j \Delta \tilde{q}_{t-j} + u_{tk} \quad (4)$$

>From regression (4) one can see that the selection of the k^{th} truncation lag is crucial. Ng and Perron (2001) show that, in the presence of a strong negative MA coefficient $\hat{\beta}$ is highly biased if the lag truncation, k , is small because u_{tk} is serially correlated. To select the optimal k , that accounts for the inverse non-linear dependence between the bias in $\hat{\beta}$ and the selected k and avoids selecting a large k when it is not needed, they propose the modified Akaike information criteria ($MAIC$). In the search procedure, the maximum starting value for k shall be data dependent and one should reset k_{\max} by a higher number and re-optimize the $MAIC$ function to confirm the optimal choice.

The modified Phillips-Perron using the GLS detrended data (MZ_α^{GLS}) is due to Ng and Perron (2001). This test requires estimation (4) with $k = 0$ and the variance and long-run variance of u_{t0} . The MZ_α^{GLS} test statistic is given by $MZ_\alpha^{GLS} = \frac{T^{-1}\tilde{q}_T^2 - s_{AR}^2}{2T^{-2}\sum_{t=1}^T \tilde{q}_{t-1}^2}$. The autoregressive estimate of the spectral density function at frequency zero of u_{t0} is given by $s_{AR}^2 = s_u^2 / \left(1 - \sum_{j=1}^k \hat{\beta}_j\right)^2$, where $s_u^2 = T^{-1} \sum_{t=k+1}^T \hat{u}_{tk}^2$, with $\hat{\beta}_j$ and $\{\hat{u}_{tk}^2\}$ obtained from equation (4), and k is chosen by the *MAIC*. Asymptotic critical values for both tests, ADF^{GLS} and MZ_α^{GLS} , are reported in Ng and Perron (2001).

The presence of structural breaks, a common feature among Latin America countries during the period, can severely bias unit root tests. Perron (1997) proposes a test that allows for a change in both intercept and slope at time T_b , which is made perfectly correlated with the data³. The test entails OLS estimation of the following innovational outlier (IO) model:

$$q_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha q_{t-1} + \sum_{j=1}^k c_j \Delta q_{t-j} + e_t \quad (5)$$

where $DU_t = 1(t > T_b)$ and $D(T_b)_t = 1(t = T_b + 1)$ with $1(\cdot)$ being the indicator function. The test is a t-statistic for $\alpha = 1$ in (5). The time T_b is chosen as $t_\alpha^* = \text{Min}_{T_b} t_\alpha(T_b, k)$, the minimum t-statistic for testing the unit root hypothesis ($\alpha = 1$). The truncation lag, k , is selected according to a *t - test general - to - specific* procedure. Critical values are found in Perron (1997).

³This tests overcomes a common criticism to Perron (1989) where the time of the break is assumed to be known a priori but, in fact, it might be correlated with the data.

A potential problem with Perron (1997) test is that it assumes no structural break under the null of unit root. Lee and Strazicich (2001) show that this assumption can result in spurious rejections when it is not true. The two-break minimum LM unit root test, due to Lee and Straizicich (2003), is unaffected by whether or not there is a break under the null. The test statistic is obtained from:

$$\Delta q_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum_{j=1}^k \gamma_j \Delta \tilde{S}_{t-j} + \varepsilon_t \quad (6)$$

where $\tilde{S}_t = q_t - \tilde{\Psi}_x - Z_t \tilde{\delta}$, $t = 2, \dots, T$; $\tilde{\delta}$ are the coefficients from the regression of Δq_t on ΔZ_t and $\tilde{\Psi}_x$ is the restricted MLE of $\Psi_x (\equiv \Psi + X_0)$ given by $q_1 - Z_1 \tilde{\delta}$. The $\Delta \tilde{S}_{t-j}$ terms are included to correct for possible serial correlation and Z_t is a vector of exogenous variables contained in the data generating process. The null of unit root is given by $\phi = 0$ and the LM test statistic, called $\tilde{\tau}$, is the t -statistic under the null. Time of the breaks ($\lambda_i = T_{Bi}/T$, $i = 1, 2$) are given by points where $\tilde{\tau}$ -statistic is at a minimum. Critical values were tabulated by Lee and Straizicich (2003).

Non-linearity in time series, which are not captured by structural changes, also leads to distortions in unit root tests. Kapetanios, Shin and Snell (2003), hereafter KSS, proposed a test to detect the presence of a unit root against nonlinear but globally stationary exponential smooth transition autoregressive (ESTAR) process. Due to identification problem under the null, KSS reparameterize the ESTAR model

and derive the test equation:

$$\Delta \bar{q}_t = \sum_{j=1}^k \rho_j \bar{q}_{t-j} + \delta \bar{q}_{t-1}^3 + \varepsilon_t \quad (7)$$

The truncation lag in (7) is meant to correct for potentially serially correlated errors. It might be selected by a *general-to-specific* approach based on the *t-test*. KSS show that the *t-statistic* under the null hypothesis $\delta = 0$ follows a non-standard distribution and provide simulated critical values.

4 Panel Data Unit Root Tests

The major reason for using panel data is that it increases power of unit root tests. We first consider panel tests for the null of unit root, as proposed by Levin, Lin, and Chu (2002), Im, Pesaran, and Shin (2003), Maddala and Wu (1999), and Taylor and Sarno (1998). The later authors suggest a multivariate ADF test based on Abuaf and Jorion (1990). Shortly, those tests are labelled as LLC, IPS, MW, and MADF, respectively. Serially correlated residuals are accounted for by including an appropriate lag truncation in each test equation.

The LLC test estimates the following regression:

$$\Delta q_{i,t} = \alpha_i + \delta q_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta q_{i,t-j} + u_{i,t} \quad (8)$$

where $i = 1, 2, \dots, N$ and $t = 1, 2, \dots, T$. The test statistic for the null of a common unit root ($\delta = 0$) is obtained from pooled regression (8) and has limiting distribution given by a $N(0, 1)$. Notice that homogeneity of δ implies that rejection of the null can occur even when only a small subset of series are stationary.

The IPS test allows for some heterogeneity in the test equation (8) by estimating individual-specific unit root coefficient δ_i . The test statistic is the sample mean of the t-statistic resulting from individual regressions estimated for each series of the panel. IPS show that the test statistic also converges to a standard normal distribution.

The MW test combines *p-values* from individual ADF regressions. Let p_i be the *p-value* for the null hypothesis that $\delta_i = 0$ in the i^{th} ADF regression. Under the null that all series in the panel have a unit root against the alternative that at least one series is stationary, the test statistic is $MW = -2 \sum_{i=1}^N \log(p_i)$, which converges to a χ_{2N}^2 . The MW test also applies to the Phillips-Perron (PP) version of the individual unit root regressions.

The MADF test estimates a multivariate version of equation (8) without the $q_{i,t-1}$ variable and with a common truncation lag ($k_i = k$). The parameters are estimated by SUR in a system of N equations. It is then conducted a jointly test on $\sum_{j=1}^k \phi_{i,j} - 1 = 0$ for all N equations of the system. The resulting Wald statistic is taken as the MADF statistics.

Changing the null hypothesis to read stationarity avoids the criticism that the null of unit root is frequently rejected if only a subset of series in the panel is stationary.

Rejection of the null of unit root in parallel to non-rejection of the null of stationary leads to the conclusion that all series in the panel are stationary.

The Nyblom and Harvey (2000), shortly NH, test is a multivariate version of the time series unit root test developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992), known as KPSS⁴. NH consider the following model with $N - vector$ time series:

$$q_t = \mu_t + \varepsilon_t, \quad \text{with } \varepsilon_t \sim N\left(0, \sum_{\varepsilon}\right) \quad (9)$$

$$\mu_t = \mu_{t-1} + \eta_t, \quad \text{with } \eta_t \sim NID\left(0, \sum_{\eta}\right), \quad t = 1, 2, \dots, T \quad (10)$$

where $q_t = (q_{1,t}, q_{2,t}, \dots, q_{N,t})'$ and $\mu_t = (\mu_{1,t}, \mu_{2,t}, \dots, \mu_{N,t})'$ is a vector random walk. NH derive the test statistic under the null hypothesis that there is no random walk in the system $\left(rank \sum_{\eta} = 0\right)$ against the alternative that at least one series is a random walk $\left(rank \sum_{\eta} > 0\right)$. Failure to reject the null indicates that the series in the panel are stationary.

As in the time series case, however, structural breaks can severely bias panel data unit root tests. To account for structural changes, we apply tests proposed by Im, Lee, and Tielsau (2005) and Papell (2002). The former is a LM test that allows for at most two structural breaks while the second allows for three structural breaks. In both tests, the time of the breaks are selected endogenously and they must coincide among the series in the panel.

⁴Hadri (2000) also proposed a test for the null of stationarity based on KPSS.

The test by Im, Lee, and Tielsau (2005) is an extension of panel LM unit root test. The test equation, which corrects for autocorrelation, is:

$$\Delta q_{i,t} = \gamma_{2,i} + \delta_i \Delta D_{i,t} + \beta_i \tilde{S}_{i,t-1} + \sum_{j=1}^{k_i} \rho_{i,j} \Delta \tilde{S}_{i,t-j} + u_{i,t} \quad (11)$$

where $\tilde{S}_{i,t-1} = q_{i,t-1} - \tilde{\gamma}_{2,i}(t-1) - \tilde{\delta}_i D_{i,t-1}$ and $\tilde{\gamma}_{2,i}$ and $\tilde{\delta}_i$ are obtained as OLS estimators in the regression $\Delta q_{i,t} = \gamma_{2,i} + \delta_i \Delta D_{i,t} + \varepsilon_{i,t}$. The dummy variable is $D_{i,t} = 1$ if $t \leq T_{B,i}$ and $D_{i,t} = 0$ otherwise. The LM statistic is the average t-statistic for $\beta_i = 0$, $i = 1, 2, \dots, N$, in regression (11). Im, Lee, and Tielsau (2005) show that the LM statistic, under the assumption that $N/T \rightarrow \kappa$ (a finite constant), converges to a $N(0, 1)$.

Papell (2002) allows for restricted structural change at three distinct dates. Restrictions impose the PPP under the alternative hypothesis. The test is a three steps procedure. Firstly, the time of the breaks are chosen by estimating SUR regressions of the form $q_{i,t} = \alpha_i + \gamma_1 D1_t + \gamma_2 D2_t + \gamma_3 D3_t + \tilde{q}_{i,t}$ subject to the PPP restrictions $\gamma_1 + \gamma_2 + \gamma_3 = 0$ and $\gamma_1 (D3 - D1) + \gamma_2 (D3 - D2) = 0$, which imposes a constant mean prior to the first and following the third break and restricts these two means to be equal. The break dates are chosen endogenously to maximize the joint log-likelihood.

Secondly, the time series are detrend according to $q_{i,t} = \alpha_i + \gamma_{i,1} D1_t + \gamma_{i,2} D2_t + \gamma_{i,3} D3_t + \tilde{q}_{i,t}$, where the estimated coefficients are allowed to vary between countries but the time of the breaks are restricted to be the same.

Finally, the t-statistic for the null $\delta = 0$ is computed in the SUR regression:

$$\Delta \tilde{q}_{i,t} = \delta \tilde{q}_{i,t-1} + \sum_{j=1}^{k_i} \phi_{i,j} \Delta \tilde{q}_{i,t-j} + \varepsilon_{i,t} \quad (12)$$

where the null hypothesis of a unit root without structural change is tested against the alternative of stationarity with PPP restricted structural change. Critical values are model-specific and computed as in Papell (2002) by Monte Carlo simulations⁵.

5 Empirical Evidence

5.1 Data Description

The data set is composed of monthly time series in the period of 1981:01 to 2003:12 for all 26 Latin-American countries: Argentina, Bahamas, Barbados, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominica, Dominican Republic, Ecuador, El Salvador, Guatemala, Haiti, Honduras, Jamaica, Mexico, Netherlands Antilles, Nicaragua, Paraguay, Peru, St. Lucia, Suriname, Trinidad and Tobago, Uruguay, Venezuela. The reference currency is the US dollar and inflation rates, for all countries, were represented by the consumer price indexes (CPI). In the empirical evidence, we consider the logarithm of the real exchange rates. All variables were obtained from the

⁵See Papell (2002), section 3.2, for details. We thank him for kindly sending us his RATS codes used to test and simulate critical values based on re-sampling bootstrapping. Given our panel dimensions, in a pentium 4 with 512 MB of ram, it took about four full days to get the simulations done.

5.2 Time Series Tests

The first set of results, which are displayed Table 1, refer to linear unit root tests without structural breaks. The traditional ADF and Z_α tests do not reject the null of unit root in most of the series. The truncation lag was selected according to the AIC. One can see that, at the standard 5% significance level, the PPP holds only for 6 countries. The high rate of rejection might be due to the widely reported lack of power of the ADF and Z_α tests.

We, then, apply the new unit root tests proposed by Elliott, Rothenberg and Stock (1996) and Ng and Perron (2001), labelled $MADF^{GLS}$ and MZ_α^{GLS} , which have better power and size properties than the traditional ones. The results reported in the last three columns of Table 1 show that, at the 5% level, the unit root is now rejected for 17 out of 26 countries. At the 10% level, the unit root is rejected for 21 countries. This is a significant improvement in the previous results. However, as discussed earlier, the presence of structural breaks might affect the performance of the new tests.

In Table 2 we take care of structural breaks by applying the tests proposed by Perron (1997) and Lee and Straizicich (2003). The first test is based in Perron (1989), while the second is a LM test. Both of them endogenously select the time of the break and the lag truncation in the test regression. The results do not add much to the

conclusions as both tests did a poor job in rejecting the null of unit root for most of the countries. Only two countries (Argentina and Trinidad and Tobago) were added to the group of 21 countries identified as having stationary real exchange rates by the $MADF^{GLS}$ and MZ_{α}^{GLS} tests. This bad performance might be due to the lack of power of the previous tests, which we overcome by applying panel data structural break tests.

Finally, we apply the nonlinear unit root test proposed by Kapetanios, Shin and Snell (2003), labelled as KSS. The results in Table 3 clearly indicate that it is not a problem of nonlinearity that lead to a unit root in the Latin-American real exchange rates. The null is reject for only one country under KSS2 and for none under KSS1. Considering the optimal lag selection, given by the MAIC, the results under KSS3 reject a unit root for 5 countries. Thus, the KSS class of tests had a poor performance in testing for the PPP hypothesis in the present sample.

5.3 Panel Data Tests

We start with the tests for the null hypothesis of a unit root. The results are reported in Table 4. For the LLC IPS MW-ADF MW-PP tests, the lag selection is country-specific and was based on the AIC. The test statistics indicate that the null is rejected at 95% (LLC and MW-ADF) and 99% (IPS and MW-PP) confidence levels. The MADF test applies the same lag value to all individuals in the panel, also selected by the AIC. It also rejected a unit root at the 99% confidence level and the result was

not sensible to alternative truncation lag. Thus, the results displayed in Table 4 are in line with the previous evidence by the new time series unit root tests and indicate that PPP holds during the period.

Given that the previous tests are sensible to the presence of a few stationary series in the panel, we apply the test by Nyblom and Harvey (2000). The result for the null of stationarity is reported in Table 5. The test uses the Newey-West bandwidth selection and the Bartlett kernel to compute the residual covariance matrix. The test statistic does not reject the null at 95% confidence level and so confirms that PPP holds in Latin-America.

Finally, we accounted for successive structural breaks that affected Latin-American real exchange rates during the period. Table 6 display the results for two structural change panel data unit root tests. We follow Im, Lee, and Tielsau (2005) and allow for a maximum of two structural breaks in each time series. The truncation lag at each possible shift is chosen according to a *general-to-specific* procedure based on the statistical significance at 10% level of the last lagged coefficient. A grid search over the interval $[0.1T, 0.9T]$ is used to determine the break locations according to the t - test on the dummy coefficients. As stressed by Im, Lee, and Tielsau (2005), the number and location of the breaks and the truncation lags are jointly determined for each unit of the panel. The test statistic reported in Table 6 provides support for the PPP, as the null of unit root is rejected at 99% of confidence.

The test proposed by Papell (2002) also reports results favorable to the PPP

hypothesis⁶. The unit root is rejected at the 99% significance level. The test allows for three common structural changes in each time series and the truncation lag, which is country specific, is chosen according to a *general-to-specific* approach based on 10% significance level of the last lagged coefficient. The dates of the breaks are endogenously chosen to maximize the joint likelihood of equation (??). The rejection of a unit root in the real exchange rates by Papell's test strengthens the evidence of PPP for Latin-American countries. The PPP holds also when it is taken into account multiple structural shifts that have long characterized those economies.

Thus, in a panel data environment, where tests are more powerful towards rejection of a unit root, one can conclude that there is strong evidence in favor of the PPP hypothesis for the Latin-American countries. This conclusion is not sensible to either changes in the null hypothesis to read stationarity instead of unit root or multiple structural changes in the individual series of the panel data.

Given the previous results of stationary real exchange rates, we computed half-lives of disturbances to PPP⁷. The results are reported in the last column of Table 3 and show that it takes on average 3 years to correct half of any PPP deviation. The pooled OLS indicates a faster convergence to the panel as whole, where the same adjustment took only 1.2 year. In general, these findings are in line with Taylor (2002) where, for a different sample of countries, the mean half-life was 2.6 years in

⁶We thank David Papell for kindly sending us his RATS codes used to perform the test and compute critical values.

⁷Following the literature, the half-life (h) is computed from an AR(1) process for the real exchange rate $q_t = \phi q_{t-1} + \varepsilon_t$ as $h = \ln(0.5)/\ln(\phi)$.

the recent floating period.

6 Concluding Remarks

This paper has performed a comprehensive analysis of the PPP hypothesis for all Latin-American countries using both time series and panel data unit root tests.

In the time series framework, we applied the traditional ADF and Phillips-Perron tests and new unit root tests, due to Elliott, Rothenberg and Stock (1996), and Ng and Perron (2001). We also allowed for structural changes and nonlinearity in real exchange rates. The results from the new tests indicated that PPP holds for the vast majority of the countries. However, structural-break and non-linear unit root tests were able to reject the null of integrated real exchange rates for a few countries. The bad performance of the structural break tests is due to their lack of power while non-linearity seems not to be a problem for Latin-American real exchange rates. To improve power of the tests, we migrated to a panel data environment.

The results of the panel data unit root tests confirmed the evidence by the new time series unit root tests. The tests for the null of a unit root unanimously indicated that real exchange rates are stationary. Due to the common criticism that these tests over-reject in the presence of few stationary series in the panel, we applied a test for the null of stationarity. The Nyblom and Harvey (2000) test confirmed the previous evidence in favor of the PPP.

Finally, we allowed for multiple structural breaks in the individual series of the

panel and applied tests proposed by Im, Lee, and Tielsau (2005) and Papell (2002). The first test allow for two breaks while the second considers up to three breaks at common dates in the time series. Both of them rejected the null of a unit root and reinforced the conclusion of stationary real exchange rates. Thus, our results show strong evidence that PPP holds for Latin-American countries in the post-1980 period.

This finding is in line with our argument that Latin-American countries share important features in their recent economic history, which must be taken into account in economic analysis. They went through debt crises, high inflation, successive economic stabilization plans, changes in exchange rate regimes, currency crises, among others. The comovement of the main economic variables associated to contagious effects of currency crises may help to explain the non rejection of the PPP hypothesis. When the Latin American countries are taken as whole in a panel, these common features and contagious effects are accounted for and our results show that the pooled real exchange rate is stationary.

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Table 1: Time Series Unit Root Tests

Countries	Traditional Tests			New Tests		
	ADF	Z_α	Lags	$MADF^{GLS}$	MZ_α^{GLS}	Lags
Argentina	-3.32**	-2.96**	4	-2.11	-0.43	14
Bolivia	-6.44***	-3.82***	14	-6.60***	-34.34***	3
Brazil	-2.04	-1.94	1	-6.36***	-32.15***	3
Chile	-1.83	-1.88	7	-2.64*	-3.43	14
Colombia	-1.44	-1.15	12	-1.74	0.82	14
Costa Rica	-4.11***	-3.75***	10	-2.81*	-4.18	12
Dominican Republic	-2.36	-1.81	0	-2.73*	-3.86	12
Ecuador	-1.69	-1.74	0	-3.20**	-5.56	12
El Salvador	-2.86*	-2.86*	0	-6.19***	-33.71***	9
Guatemala	-2.14	-2.86*	0	-5.93***	-30.41***	3
Haiti	-1.98	-1.86	14	-2.96**	-4.36	12
Honduras	-1.72	-1.64	2	-6.18***	-33.22***	3
Mexico	-2.85*	-2.84*	10	-3.45***	-7.58	12
Nicaragua	-3.37**	-3.23**	14	-5.91***	-30.08***	14
Paraguay	-2.29	-1.22	0	-6.06***	-29.77***	3
Peru	-2.28	-2.57*	12	-6.65***	-38.04***	16
Uruguay	-1.70	-1.64	0	-6.09***	-29.35***	3
Venezuela	-2.34	-2.07	2	-1.66	0.11	14
Bahamas, The	-3.13**	-1.25	13	-2.99**	-3.19	12
Barbados	-2.50	-1.21	12	-2.11	-1.15	14
Dominica	-2.67*	-0.81	0	-5.84***	-28.56***	3
Jamaica	-1.83	-2.00	0	-6.14***	-31.54***	3
Netherlands Antilles	-1.53	-0.79	0	-6.01***	-29.31***	3
St. Lucia	-2.01	-0.99	2	-6.47***	-35.87***	3
Suriname	-3.15**	-2.67*	0	-3.64***	-10.19	12
Trinidad and Tobago	-2.24	-1.34	0	-2.11	-0.42	14
Critical Values						
1%	-3.45	-3.45		-3.42	-23.80	
5%	-2.87	-2.87		-2.91	-17.30	
10%	-2.57	-2.57		-2.62	-14.20	

Notes: *** - the unit root is rejected at 99% confidence level. ** - the unit root is rejected at 95% confidence level. * - the unit root is rejected at 90% confidence level.

Table 2: Time Series Unit Root Tests under Structural Breaks

Countries	Perron (1997)				Lee and Strazicich (2003)				
	Stat.	T_b	Lags	CV(5%)	Stat.	T_{b1}	T_{b2}	Lags	CV(5%)
Argentina	-5.44*	90-01	10	-4.80	-5.06	91-04	01-08	10	-5.29
Bolivia	-8.61*	87-07	11	-4.80	-5.91*	90-02	91-11	7	-5.29
Brazil	-3.40	98-10	12	-4.80	-4.83	86-09	96-11	12	-5.29
Chile	-5.89*	85-05	11	-4.80	-4.84	86-09	96-11	12	-5.29
Colombia	-3.73	85-01	2	-4.80	-4.52	83-09	96-06	12	-5.29
Costa Rica	-6.77*	92-03	10	-5.08	-4.16	86-10	96-07	11	-5.29
Dominican Republic	-7.98*	84-11	0	-5.08	-3.73	85-09	97-04	12	-5.29
Ecuador	-5.09*	85-10	10	-4.80	-7.78*	93-07	94-06	3	-5.29
El Salvador	-6.54*	85-11	1	-5.08	-4.59	85-09	93-02	12	-5.29
Guatemala	-13.07*	86-04	0	-5.08	-7.81*	84-07	85-11	10	-5.29
Haiti	-4.44	96-02	12	-4.80	-4.42	88-11	99-03	12	-5.29
Honduras	-15.10*	90-02	11	-5.08	-4.54	86-06	97-08	11	-5.29
Mexico	-4.75	85-05	10	-4.80	-3.83	85-02	95-03	12	-5.29
Nicaragua	-10.15*	87-12	12	-5.08	-5.73*	85-02	95-07	10	-3.84
Paraguay	-3.83*	83-12	0	-4.80	-4.03*	91-02	91-02	12	-3.84
Peru	-5.23*	89-06	12	-5.08	-3.50	85-11	88-07	10	-3.84
Uruguay	-3.43	02-06	12	-4.80	-4.52	85-11	86-06	1	-5.29
Venezuela	-4.39	86-10	2	-5.08	-5.75*	86-04	87-11	9	-5.29
Bahamas, The	-4.05	86-10	12	-4.80	-3.27	91-08	94-08	12	-3.84
Barbados	-4.44	96-11	12	-4.80	-3.25	85-06	95-03	10	-3.84
Dominica	-4.38	85-05	12	-4.80	-8.71*	87-02	88-06	10	-5.29
Jamaica	-4.71	83-09	11	-4.80	-4.88	89-07	98-10	12	-5.29
Netherlands Antilles	-4.01	85-05	12	-4.80	-5.91*	87-11	90-09	2	-5.29
St. Lucia	-4.02	85-05	12	-4.80	-4.12	87-11	99-04	12	-5.29
Suriname	-13.97*	94-04	4	-5.08	-5.16	89-03	01-06	0	-5.29
Trinidad and Tobago	-5.81*	85-09	0	-5.08	-4.39	86-10	00-02	12	-5.29

Notes: * - the unit root is rejected at 95% confidence level.

Table 3: Nonlinear Unit Root Test and Half-lives PPP

Countries	$KSS1$	$KSS2$	$KSS3$	Half-lives
Argentina	-0.80	-1.21	-2.87*	0.8
Bolivia	-0.97	-1.87	-5.16***	0.4
Brazil	-0.80	-2.01	-1.50	2.1
Chile	0.93	-2.12	-2.62	3.6
Colombia	1.59	-2.51	-2.25	9.6
Costa Rica	0.05	-2.00	-1.58	0.6
Dominican Republic	-0.30	-0.37	-0.91	2.1
Ecuador	0.26	-2.36	-2.30	3.6
El Salvador	-1.49	-1.35	-1.21	1.2
Guatemala	-0.60	-0.19	-0.43	2.2
Haiti	-0.86	-2.00	-2.37	2.0
Honduras	-0.21	0.60	-0.98	1.6
Mexico	-0.84	-2.04	-4.02***	1.1
Nicaragua	-0.40	-1.73	-4.06***	0.8
Paraguay	1.11	-1.32	-1.29	5.2
Peru	-0.14	-1.55	-0.96	1.3
Uruguay	-1.27	-0.64	-1.67	2.5
Venezuela	-0.69	-2.32	-2.41	1.5
Bahamas, The	-0.69	-1.98	-1.16	0.9
Barbados	0.00	-1.92	-2.07	4.1
Dominica	0.50	-1.58	-1.14	6.4
Jamaica	-0.31	-1.99	-1.70	2.6
Netherlands Antilles	0.66	-2.19	-1.24	11.5
St. Lucia	0.46	-1.91	-1.18	5.2
Suriname	-1.65	-2.92**	-3.44**	1.1
Trinidad and Tobago	-0.41	-1.78	-1.62	4.4
Critical Values			Mean	3.0
1%	-3.48		Med.	2.1
5%	-2.93		SD	2.8
10%	-2.66		Pooled	1.2

Notes: *** - the unit root is rejected at 99% confidence level. ** - the unit root is rejected at 95% confidence level. * - the unit root is rejected at 90% confidence level. $KSS1$ has $k = 0$ in equation (7). In $KSS2$, k is chosen by a *general-to-specific* procedure based on the *t-test*. In $KSS3$, k is chosen by the optimal *MAIC*.

Table 4: Panel Data Unit Root Tests for the Null of Unit Root

LLC	IPS	MW-ADF	MW-PP	MADF	lags
-1.72**	-2.48***	70.13**	83.44***	130.53***	4

Notes: *** - the null of unit root is rejected at 99% confidence level. ** - the null of unit root is rejected at 95% confidence level. The lag selection for the LLC, IPS, MW-ADF, and MW-PP is country-specific and was based on the AIC. For the MADF, alternative lag values were considered and the result did not change.

Table 5: Panel Data Unit Root Test for the Null of Stationarity

Nyblom and Harvey	lags
5.23	5

Notes: The 5% critical value for the NH test is 5.64. The test uses the Newey-West bandwidth selection and the Bartlett kernel to compute the residual covariance matrix.

Table 6: Panel Data Unit Root Tests under Structural Breaks

	test statistic	TB1	TB2	TB3	lags	5% CV	1% CV
Im, Lee, and Tieslau (2005)	-18.6***	86:10	93:02	-	12	-1.96	-2.58
Papell (2002)	-12.95***	85:06	85:08	90:08	CS	-10.58	-12.09

Notes: *** indicates that the null of unit root is rejected at 99% confidence level. CS - means that the lag selection is country specific [see equation (12)].